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Sensitivity of winter tourism to temperature increases over the last decades

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ABSTRACT

This paper provides new evidence on the impact of temperatures on tourism demand in the winter season. The analysis is based on time series data spanning from 1960 to 2015 for the South Tyrolean mountains in Italy. Since 1960, winter temperatures have increased by 0.4 degrees Celsius per decade, on average. A nonlinear autoregressive distributed lag (ARDL) model is employed for the estimations. This model allows two separate coefficients to impact tourism demand, following temperature changes (decreases or increases). Results reveal that an increase in winter temperatures by one degree Celsius leads to a decline in the number of accommodation guests (arrivals) by eight per cent, while temperature decreases have no effect on the number of arrivals. However, sensitivity to temperature increases has been declining since the early 1990s, probably due to the widespread usage of snowmaking facilities. The number of these facilities has increased by almost 10 per cent per year on average over the same period. In recent years (1986–2015), and as a consequence of these measures, temperature increases no longer have an effect on winter tourism demand. Conversely, decreases in temperatures lead to small increases in arrivals (by four per cent increase due to a one degree Celsius decrease).

1. Introduction

As global warming is predicted to continue throughout this century, many scholars have raised concerns about potential negative consequences on the global economy. Literature demonstrates that climate change has significant negative effects on the output and productivity of outdoor work-intensive sectors such, as agriculture (Deschênes and Greenstone, 2007; Schlenker and Roberts, 2009) or construction (Heal and Park, 2016). In addition, there is research revealing that not only weather-dependent industries are affected but also that there are wider effects of climate change on aspects such areas as economic growth (Dell, Jones and Olken, 2012), economic production (Burke et al., 2015) human health, conflict, migration, and demographics (for surveys of the literature see Dell, Jones and Olken, 2014; Carleton and Hsiang, 2016; Heal and Park, 2016). However, the relationship between climate variability and tourism demand has been largely overlooked in the economics literature (see for an exception Agiormirgianakis, Serenis and Tsounis, 2017). In contrast, climate factors and weather fluctuations are well covered by recent tourism literature, often through use of dynamic time series or panel data models (Falk, 2010; Li, Song and Li, 2016; Goh, 2012; Li, Song and Li, 2016; Töglhofer, Eigner and Prettenthaler, 2011; Zhang and Kulendran, 2017).

Independent of the field, the literature mentioned above agrees to the extent that there is a symmetric long-run relationship between year-to-year fluctuations in weather indicators or climate factors and economic outcomes. Implicitly this means that the effects of weather fluctuations such as temperature increases or decreases on outcomes are assumed to be of the same magnitude independent of direction.

The aim of the study is to analyse the impact of temperatures on demand for winter tourism measured as registered guest arrivals. The analysis is based on time series data for the province of South Tyrol (Alto Adige in Italy) for the winter period 1959/1960 to 2014/2015. This time span of more than 50 years allows for testing the stability of the regression relationship between temperatures and winter tourism demand over time. We employ a nonlinear autoregressive distributed lag (ARDL) model introduced by Shin et al. (2013), which makes it possible to capture asymmetric effects in the long-run relationships. Non-symmetric reactions could well be expected in the case of winter tourism: On the one hand, temperature increases may have negative effects on winter tourism demand because it can shorten the snow season with a reduction in snow depth and also may reduce the number of optimal snowmaking days. On the other hand, unusually low temperatures may not necessarily stimulate tourism demand further even if smaller decreases are beneficial for winter sports.

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This article contributes to existing tourism literature in several ways. First, it allows for asymmetric long-run elasticities of tourism demand. To our knowledge, the nonlinear ARDL model with bounds test for small samples developed by Narayan (2005) has not been applied to tourism demand studies. Second, in order to study consequences of a structural break caused by increased investment in snowmaking facilities over time, this paper provides estimations for different sub-periods to deal with possible impacts due to structural breaks. Third, we use temperatures to measure year-to-year fluctuations in weather. While the majority of tourism studies for winter destinations focus on the role of natural snow, little is known about the relationship between winter tourism demand and long term temperature changes, which, for instance, is the central variable in climate change scenarios reported by the Intergovernmental Panel on Climate Change (IPCC, 2014).

The structure of this paper is as follows. Section two outlines the conceptual background and empirical model together with the data sources, while section three presents the study area and descriptive statistics. Section four presents the empirical results, and section five provides concluding remarks.

2. Conceptual background and empirical model

In this study, tourism inflows are related to the year-to-year fluctuations of local weather conditions. Similar approaches are commonly used to investigate climate change impacts on agriculture (Deschênes and Greenstone, 2007; Schlenker and Roberts, 2009), electricity consumption and energy demand (Considine, 2000; Bessec and Fouquau, 2008; Do et al., 2016), and tourism demand (Falk, 2010, 2013). The underlying assumption is that weather conditions are exogenous and generally random (Deschênes and Greenstone, 2007), implying that their inter-annual variations can be used to identify the tourism impact of climate change. Commonly used measures in these cases are, for instance, growing degree days in agriculture (Schlenker and Roberts, 2009) or heating and cooling degree days in electricity consumption (Bessec and Fouquau, 2008).[1] Typically, these studies find a non-linear relationship between temperature or degree days and economic outcomes. In the case of winter tourism demand, often used indicators are snow conditions measured as snow depth or temperatures (Töglhofer et al., 2011). As an alternative, the number of snow-making days calculated as wet bulb temperatures can be used (Demiroglu et al., 2015).

According to the tourism literature, weather conditions can have a major impact on tourism inflows (Rosselló-Nadal et al., 2011; Töglhofer et al., 2011). Goméz-Martín (2005) suggests that a season with unfavourable weather conditions may prevent tourists from travelling to that region in the following years. This implies that there is a lagged response among tourists to favourable or unfavourable weather conditions. Studies using dynamic panel data models and annual data for the total winter season show that natural snow conditions significantly affect the number of overnight stays (Damm et al., 2017; Falk, 2010, 2013; Töglhofer et al., 2011), although the magnitude of the relationship is relatively modest and varies across areas and time spans.[2]

In this study, temperature is the main weather variable. The reason for this is that temperature can indicate both availability of natural snow and prevalence of suitable conditions for snow-making. In contrast to the northern side, the south of the European Alps is characterised by low natural snow cover (Marty, 2008), meaning that cold temperatures that allow for snow production are particularly important in this region. In general, temperature can have direct as well as indirect effects on winter tourism demand. Higher temperatures reduce the likelihood that sufficient snow (whether natural or fabricated) will be available, can shorten the snow season, lead to a deterioration of the snow quality, and thereby affect tourism demand. Englin and Moeltner (2004) show that lower daily temperatures lead to an increase in skier visits. Warmer weather also makes other competing outdoor activities such as hiking and biking more attractive, thereby shifting tourism demand away from the mountain destinations to other destinations. However, colder weather is not necessarily enough to stimulate tourism demand further. There is also a temperature threshold below which snow sports are no longer comfortable to perform.

Given these considerations, it is assumed that the relationship between temperature and winter tourism demand is asymmetric. Asymmetric relationships are common features of many economic relationships and are found in export demand behaviour, interest rate spreads, inflation, price dynamics, and visitors attendance to events (Katrakilidis and Trachanas, 2012; Shahzad et al., 2017; Süssmuth and Woitek, 2013; Van Hoang et al., 2016; Verheyen, 2013). In the case of tourism demand, it is likely that guest arrivals in the winter seasons react more strongly to increases than to decreases in temperature. The reaction of guest arrivals to real income might also be affected by fluctuations in the business cycle differently.

In addition, a number of factors could affect the stability of the relationship between year-to-year variations in weather and economic outcomes. Literature suggests that adaptation measures have an immense potential to reduce negative impacts of climate change (Burke et al., 2016). Among winter sport destination and ski lift companies, the most important adaptation strategies are the use of snowmaking facilities and diversification (i.e., supply of leisure activities that are not dependent on natural snow conditions or favourable temperatures). Nowadays, between 70 and 80 per cent of the ski area is covered by snowmaking facilities according to the environmental agency of South Tyrol (Agenzia provinciale per l'ambiente).[3] Although snowmaking is dependent on cool temperatures and temperatures have increased over time, the new generation of improved snowmaking technologies makes it possible to produce snow in a very short period of time. Thus, the massive investments in snowmaking facilities and technological progress may weaken the connection between temperature changes and tourism inflows over time. In fact, using a time span of 35 years, Töglhofer et al. (2011) find that the importance of the relationship between tourism demand and snow depth in winter sport destinations is decreasing over time. The authors attribute this occurrence to the widespread usage of snowmaking facilities. Being that in South Tyrol ski lift companies started to heavily invest in snowmaking facilities after the extremely mild and snow poor winter seasons of 1988/1889 and 1989/1990, analysis will be split into two sub-periods (before and after the period with massive investments in snowmaking).

The empirical model can be derived from the theory of tourism demand. Accordingly, quantity of tourism goods or services demanded is a function of income of consumers and its price. The tourism demand literature suggests that tourism inflows depend on real income and tourism prices of the origin and home country, exchange rates, and transportation costs (Peng et al., 2015; Song et al., 2009). Theory predicts that tourism demand increases as real income, and decreases if relative prices change in an unfavourable way (i.e. prices in the host (tourist) country increase relative to prices in other countries). According to Song et al. (2009), tourist arrivals are the most common measure of tourism demand. Other measures are tourist expenditure and overnight stays. The latter has the advantage that it accounts for both tourist arrivals and the length of stays. In principle, overnight stays can be regarded as the more accurate measure because they account for the length which generally tends to decline over time. Based on our sample, the two measures of tourism demand are highly correlated.[4] Therefore, we follow suit with the literature and use tourist arrivals as the dependent variable.

[1] Growing (heatting or cooling) degree days mean the number of days in which the temperatures are above or below a given threshold.

[2] Studies using skier visits (or skier days or lift transports) as the dependent variable are not considered here.

[3] http://ambiente.provincia.bz.it/contatti.asp.

[4] The correlation coefficient between the annual percentage change in arrivals and overnight stays is 0.89 and significant at the 1 per cent level based on the full sample from 1959/1960–2014/2015.

Tourism demand not only depends on economic factors but also on non-economic ones, including climate conditions and variability measured as tourism climate index or climate variability measured as weather factors (Goh, 2012; Rosselló-Nadal, 2014). We assume a log-linear tourism demand model that is augmented by average temperatures (measured in levels) in the winter season. The symmetric static long-run tourism demand model is modelled as follows (sub index t suppressed):

$$\ln ARR = \alpha + L_{GDPPC}\ln GDPPC + L_{EX}\ln EX + L_{TEMP}TEMP + v \quad (1)$$

where ln is the natural logarithm, and *ARR* is guest arrivals in accommodation establishments to the province South Tyrol/Alto Adige (Italy) for the winter season December to April (ASTAT, 2016). *GDPPC* is the real gross domestic product per capita (in purchasing power parities) weighted by the share of guest arrivals of the visitor countries. Real GDP per capita (measured in 1990 International GK$) is drawn from the Maddison database for the period 1959–2010 (Maddison project 2013) and updated using real GDP per capita data (measured in PPP constant 2011 international $) drawn from the World Development Indicators Database (World Development Indicators, 2013).[5] Weighted GDP per capita is constructed using the share of arrivals in total arrivals.

EX is the weighted real exchange rate measured as purchasing power parity of the main visitor countries relative to the destination and is drawn from the Penn world tables.[6] Seetaram et al. (2016) show that relative purchasing power parities are a better predictor of tourism demand than relative exchange rates. Note that the empirical model only accounts for relative prices and thus neglects the impact of domestic prices. It would be preferable to account for a price index for domestic visitors. However, regional consumer price indices of the destination are not available. Since more than 70 per cent of visitors are from abroad, the omission of the domestic price index is unlikely to have a strong impact on overall tourism demand.

TEMP denotes the mean of average maximum and average minimum temperatures at the monthly level for the period December to March for a representative mountain weather station. April temperatures are not relevant because some resorts are closed at the beginning or middle of April. Temperature data at the monthly level are drawn from the public weather service of South Tyrol (Servizio Meteorologico, 2017). In particular, time series on average maximum and average minimum temperatures at the monthly level are available for about 20 weather stations, of which one half are located in mountainous regions. We choose the weather station of Sesto/Sexten at an altitude of 1300 m in the middle of the ski area. We assume that the weather station is representative of other mountain stations in South Tyrol. Note that 1300 m is the typical elevation of the valley lift station of many ski areas in that region. Other weather stations have gaps such as Dobbiacco/Toblach or start in the 1980s or 1990s (Selva/Wolkenstein).[7] It is important to note that official data on snow cover (such as snow depth) for weather stations in South Tryol (Alto Adige) are only available from 1980 onwards.[8]

The tourism demand model (1) characterises to what extent tourism arrivals are affected by changes in the real GDP of the visitor countries, real exchange rates and temperatures in the winter season. We expect the long-run coefficient, L_{GDPPC}, i.e. the income elasticity of arrivals, to be positive. Thus, increases in real income of the visitor countries lead to an increase in arrivals in South Tyrol. Furthermore, an increase in EX, i.e. the real purchasing power, reflects an increase in the purchasing power of foreign currencies. In turn, this results in an increase in arrivals, with the long-run coefficient L_{EX} expected to be positive as well. For the impacts of temperature, it is obvious that the higher the average temperatures are, the lower the tourism demand in the winter season is. One reason for the negative relationship is that higher temperatures decrease the number of optimal days for snowmaking production. The other reason is that winter tourists prefer cold snow, and warmer temperatures make other outdoor activities in the winter season (hiking in the snow, snowshoeing) more attractive. Thus, we expect a negative long-run coefficient of L_{TEMP}.

Equation (1) represents the long-run relationship between arrivals, income, relative prices and temperatures. It is assumed implicitly that the relationships are symmetric following increases or decreases in temperatures or in the economic determinants such as real GDP and relative prices.

The tourism demand model can be specified as an ARDL model which is introduced in tourism demand modelling by Narayan (2004) and suggested in the literature on tourism impacts of climate change and climate variability (Rosselló-Nadal, 2014).

$$\begin{aligned}\Delta\ln ARR_t = {} & \alpha_0 + \alpha_1 lnARR_{t-1} + \alpha_{GDPPC}\ln GDPPC_{t-1} + \alpha_{EX} lnEX_{t-1} \\ & + \alpha_{TEMP}TEMP_{t-1} + \sum_{i=1}^{p}\beta_i \Delta \ln ARR_{t-i} \\ & + \sum_{i=0}^{q1}\beta_i^{GDPPC}\Delta\ln GDPPC_{t-i} + \sum_{i=0}^{q2}\beta_i^{EX}\Delta\ln EX_{t-i} \\ & + \sum_{i=0}^{q3}\beta_i^{TEMP}\Delta TEMP_{t=i} + u_t \end{aligned} \quad (2)$$

The least square estimator can be used to estimate the tourism demand model. The optimal lag lengths for p, $q1$, $q2$, and $q3$ can be determined by R-squares, and/or standard Akaike, and/or Schwarz and/or Hannan-Quinn information criteria. But more importantly, the determination of these parameters should take into consideration the autocorrelation of the residuals of u. In other words, if we detect series correlated u, we need to adjust these parameters (usually by increasing them) to eliminate the autocorrelation. At this stage, we shall also carry out more diagnostic tests: the Jarque-Bera normality, the Breusch-Pagan's heteroscedasticity, the Ramsey functional form, autoregressive conditional heteroskedasticity (ARCH), and the cumulative sum of the recursive residual (CUSUM) and CUSUM square (CUSUMSQ) for stability of parameters. The latter two are tests based on the criteria of the 95% confidence band (see Fig. 3 in the Appendix for the diagnostic tests).

The existence of the long-run relation (1) is identified by the bounds test. The null of no cointegration is characterised as

$$H_0: \ \alpha_1 = \alpha_{GDP} = \alpha_{EX} = \alpha_{TEMP} = 0 \quad (3)$$

Note that distribution of the error term of equation (2) is not standard when some of the variables involved are I(1) processes. Pesaran et al. (2001) therefore propose a bounds-test. If the *F*-value from (3) is above the upper bound, we can reject the null hypothesis of (3). Narayan (2005) notes that the critical values provided in Pesaran et al. (2001) are only valid for a relatively large sample size. For smaller samples, Narayan (2005) provides the critical values to carry out a bounds test.

The ARDL model of (2) can be expressed in terms of an error-correction model (ECM):

$$\begin{aligned}\Delta\ln ARR_t = {} & \alpha_0 + \alpha_1(lnARR_{t-1} + L_{GDPPC}\ln GDPPC_{t-1} + L_{EX} lnEX_{t-1} \\ & + L_{TEMP}TEMP_{t-1}) + \sum_{i=1}^{p-1}\beta_i\Delta\ln ARR_{t-i} + \sum_{i=0}^{q1-1}\beta_i^{GDP}\Delta\ln GDP_{t-i} \\ & + \sum_{i=0}^{q2-1}\beta_i^{EX}\Delta\ln EX_{t-i} \sum_{i=0}^{q3-1}\beta_i^{TEMP}\Delta TEMP_{t-i} + u_t \end{aligned} \quad (4)$$

The long-run coefficients are obtained from the following relations:

$$\alpha = -\alpha_0/\alpha_1$$

$$L_{GDPPC} = -\alpha_{GDPPC}/\alpha_1$$

[5] The break in GDP per capita due to German unification is overbridged using change in real GDP per capita for West Germany in 1990.

[6] http://www.rug.nl/ggdc/productivity/pwt/.

[7] In general, temperature data of the various weather stations are highly correlated (with correlation coefficients between 0.90 and 0.98).

[8] We would like to thank one of the reviewers and Philipp Tartarotti from the public weather service institution (Servizio Meteorologico) for this information.

$$L_{EX} = -\alpha_{EX}/\alpha_1$$

$$L_{TEMP} = -\alpha_{TEMP}/\alpha_1$$

The assumption that the relationships are symmetric – implying that increases or decreases in temperatures or in the economic determinants real GDP and relative prices have the same impact – is problematic. High temperature can have negative effects on arrivals. Tourists are less interested in winter sport activities when it is unusually warm. Warm temperatures lead to a deterioration of the quality of snow and also limit the opportunity to produce snow. Colder temperatures can stimulate tourism demand in the winter season. However, this is unlikely to have been the case in the earlier decades with phases of very cold winter temperatures. Similarly, the reaction of tourists to positive income shocks might be different from the response to negative shocks. Furthermore, tourism effects of depreciations can be higher in absolute terms than those of appreciations. However, relative prices are likely to be less relevant after Italy joined the European monetary zone in the mid-1990s.

In order to distinguish between the effects of a temperature increase and decrease, we use the nonlinear ARDL framework introduced by Shin et al. (2013). As the ARDL, a nonlinear ARDL approach can be used when explanatory variables are I(1) and I(0) or a combination of both (Pesaran et al., 2001). The long-run symmetries are tested using the Wald test.

To allow for asymmetric impacts, we extend the discussion above based on their corresponding partial sums of positive and negative changes:

$$\ln GDPPC_t^+ = \sum_{i=1}^{t} \max(\Delta \ln GDPPC_i, 0)$$

$$\ln GDPPC_t^- = \sum_{i=1}^{t} \min(\Delta \ln GDPPC_i, 0)$$

$$\ln EX_t^+ = \sum_{i=1}^{t} \max(\Delta \ln EX_i, 0)$$

$$\ln EX_t^- = \sum_{i=1}^{t} \min(\Delta \ln EX_i, 0)$$

$$TEMP_t^+ = \sum_{i=1}^{t} \max(\Delta TEMP_i, 0)$$

$$TEMP_t^- = \sum_{i=1}^{t} \min(\Delta TEMP_i, 0)$$

Since only positive changes will be added into positive partial sums, $\ln GDPPC_t^+, \ln EX_t^+$, and $TEMP_t^+$ are monotonically increasing functions. In contrast, for the same reason, $\ln GDPPC_t^-, \ln EX_t^-$, and $TEMP_t^-$ are monotonically decreasing functions. Thus, the original series can be decomposed into the sum of these corresponding two trends: $\ln GDPPC_t = \ln GDPPC_0 + \ln GDPPC_t^+ + \ln GDPPC_t^-$, $\ln EX_t = \ln EX_0 + \ln EX_t^+ + \ln EX_t^-$, and $TEMP_t = TEMP_0 + TEMP_t^+ + TEMP_t^-$.

The long-run relation is now formulated as follows (sub index t suppressed):

$$\ln ARR = \alpha + L_{GDPPC}^+ \ln GDPPC^+ + L_{GDPPC}^- \ln GDPPC^- + L_{EX}^+ \ln EX^+ + L_{EX}^- \ln EX^- + L_{TEMP}^+ TEMP^+ + L_{TEMP}^- TEMP^- + v \quad (5)$$

The specification of (5) allows us to separately study impacts due to positive and negative changes. The asymmetry is not only characterised by differences of the two long-run coefficients L_i^+ and L_i^- but also represented by various degrees of significance. For instance, the asymmetric impacts due to temperature can be expected by a significant L_{TEMP}^+ but insignificant L_{TEMP}^-.

Similar to (2), (5) can be estimated by the following:

$$\begin{aligned}\Delta \ln ARR_t = {} & \alpha_0 + \alpha_1 \ln ARR_{t-1} + \alpha_{GDPPC}^+ \ln GDPPC_{t-1}^+ \\ & + \alpha_{GDPPC}^- \ln GDPPC_{t-1}^- + \alpha_{EX}^+ \ln EXDP_{t-1}^+ + \alpha_{EX}^- \ln EX_{t-1}^- \\ & + \alpha_{TEMP}^+ TEMP_{t-1}^+ + \alpha_{TEMP}^- TEMP_{t-1}^- + \sum_{i=1}^{p-1} \beta_i \Delta \ln ARR_{t-i} \\ & + \sum_{i=0}^{(q1+)} \beta_i^{GDP+} \Delta \ln GDP_{t-i}^+ + \sum_{i=0}^{(q1-)} \beta_i^{GDP-} \Delta \ln GDP_{t-i}^- \\ & \sum_{i=0}^{(q2+)} \beta_i^{EX+} \Delta \ln EX_{t-i}^+ + \sum_{i=0}^{(q2-)} \beta_i^{EX-} \Delta \ln EX_{t-i}^- + \\ & \sum_{i=0}^{(q3+)} \beta_i^{TEMP+} \Delta TEMP_{t-i}^+ + \sum_{i=0}^{(q3-)} \beta_i^{TEMP-} \Delta TEMP_{t-i}^- + u_t \end{aligned} \quad (6)$$

Existence of the long-run relation of (4) can be tested according to the bounds test. The null hypothesis is now

$$H_0 : \alpha_1 = \alpha_{GDP}^+ = \alpha_{GDP}^- = \alpha_{EX}^+ = \alpha_{EX}^- = \alpha_{TEMP}^+ = \alpha_{TEMP}^- = 0 \quad (7)$$

The decision rule is the same as the symmetric ARDL model: if the *F*-value from the test of (6) is above the upper bound, we may reject the null of no-cointegration (7). Shin, Yu, and Greenwood-Nimmo (2013) note one interesting feature of the nonlinear ARDL model, i.e. according to the setting of asymmetric long-run relation (5), we have double as many independent variables in comparison to (1), and it becomes easier to reject the null of no-cointegration, since the upper bounds are sensitive to the number of independent variables involved in the estimation model. As pointed out by Shin et al. (2013), partial sums should not be treated as independent variables. Thus, the number of independent variables used for a bounds test in the nonlinear ARDL model should be between three and six in our case.

Similarly, the long-run coefficients in (4) can be obtained according to those estimated in (5):

$$\alpha = -\alpha_0/\alpha_1$$

$$L_{GDP}^+ = -\alpha_{GDP}^+/\alpha_1$$

$$L_{GDP}^- = -\alpha_{GDP}^-/\alpha_1$$

$$L_{EX}^+ = -\alpha_{EX}^+/\alpha_1$$

$$L_{EX}^- = -\alpha_{EX}^-/\alpha_1$$

$$L_{TEMP}^+ = -\alpha_{TEMP}^+/\alpha_1$$

$$L_{TEMP}^- = -\alpha_{TEMP}^-/\alpha_1$$

The long-run asymmetries can be tested by employing the Wald tests on

$$L_{GDP}^+ = L_{GDP}^- \; L_{EX}^+ = L_{EX}^-$$

$$L_{TEMP}^+ = L_{TEMP}^-$$

respectively. Similarly, the short-run asymmetries can be tested by the Wald tests on

$$\sum_{i=0}^{(q1+)} \beta_i^{GDP+} = \sum_{i=0}^{(q1-)} \beta_i^{GDP-}$$

$$\sum_{i=0}^{(q2+)} \beta_i^{EX+} = \sum_{i=0}^{(q2-)} \beta_i^{EX-}$$

$$\sum_{i=0}^{(q3+)} \beta_i^{TEMP+} = \sum_{i=0}^{(q3-)} \beta_i^{TEMP-}$$

respectively. Note that the Wald test statistic follows a Chi-square distribution, according to Shin et al. (2013).

In order to deal with possible structural change caused by introducing the snowmaking facilities in the nonlinear ARDL model (5), we establish two subsamples with equal size. Due to a significant reduction of the number of observations in each subsample, a semi-nonlinear ARDL is used which allows only for asymmetries in temperatures.

3. Study area

The study area is located in the province of South Tyrol/Alto Adige in Northern Italy. South Tyrol is an ideal location to study because of its long tourism tradition and the relative homogeneity of the guests. Downhill skiing, snowboarding, and cross-country skiing are the preferred leisure activities during the winter season (Franch et al., 2006). There are more than 60 ski resorts with about 370 ski lifts. The total lift capacity is 520,000 persons per hour with the total number of passengers transported uphill at about 120 million (Provincia Autonoma di Bolzano-Alto Adige Ufficio trasporti funiviari, 2015). Other popular activities, besides skiing, include relaxing, enjoyment of nature and surroundings, and hiking in the snow (Franch et al., 2006). The Eastern part of South Tyrol/Alto Adige belongs to the Dolomites, which are listed as a UNESCO world heritage site.

The analysis focuses on the winter season spanning the months December to April. November is excluded because it belongs to the non-snow holiday season and accounts for a small share of guest arrivals (five per cent of the total winter season). A breakdown by travel motivation or travel purpose is not available (leisure versus business tourists). However, given the dominance of rural and mountain destinations and the absence of larger cities (except the capital city Bolzano), the share of business tourists and cultural tourists is very low.

The ongoing global warming can be clearly observed in the study area. Descriptive statistics show that the mean of average maximum and minimum temperatures in the winter months December to March increases by 0.4 °C per decade. This trend increase is significant at the one per cent level (see Fig. 1 in the Appendix). The rate of increase is consistent with that reported for the winter months in the U.S. for a similar time period (Burakowski et al., 2008).

In South Tyrol, during the winter season 2014/2015 (defined as December to April) there are about 2.4 million arrivals, representing 10.6 million overnight stays. Average length of stay in the winter season has declined from almost 7 days in the 1960s to 4.5 in the last years. The winter season accounts for one third of total arrivals on average over the sample period. Germany has the largest share in total arrivals with about 60 per cent, followed by domestic visitors with 30 per cent (measured as sample averages) (ASTAT, 2016). Arrivals of visitors from Austria, Switzerland, the Netherlands, Belgium, the United Kingdom and the United States account for between one and four per cent. Polish and Czech visitor arrivals account for one per cent each.

Figs. 1 and 2 show the evolution of tourist arrivals, real GDP per capita and temperatures of the sample period. As can be seen, the extremely warm winter seasons 1988/1989 and 1989/1990 are associated with strong declines of about 30 and 27 per cent respectively in the number of arrivals compared to the base period of the three seasons before the extreme seasons. However, the recent temperature shock in 2006/2007 is not associated with a decline in arrivals. This already indicates that the sensitivity of winter tourism demand with respect to temperature is declining over time.

4. Empirical results

We first analyse the order of integration for the time series. ADF tests show that guest arrivals, GDP per capita in constant prices, and the weighted relative purchasing power parity are all integrated of order one while temperature is stationary. Arrivals and real GDP per capita became stationary at first differences. We report results obtained from both the ARDL and nonlinear ARDL approach. The long-run coefficients of the ARDL models assuming symmetric reaction show that there is a significant long run relationship between arrivals, weighted GDP per capita, relative purchasing power parity and temperatures. The income elasticity is about 2.4 and significantly higher than one, indicating that arrivals disproportionally increase when real GDP per capita increases by one per cent.

The real exchange rate variable measured as relative purchasing power parity has the expected positive sign and is significant at the five per cent level. The coefficient of 0.77 indicates that a real depreciation of the domestic currency by 1 per cent leads to an increase in total arrivals by 0.8 per cent. Temperatures in the winter season are significantly negative, indicating that warmer weather leads to a lower number of arrivals. An increase in temperature of 1 °C leads to a decline in the number of tourist arrivals by 4.6 per cent. Further, it should be noted that the ARDL model passes all diagnostic tests in our considerations except the CUSUMSQ test. This indicates that the model is subjected to non-stable variance of residuals, implying non-stable coefficients.

The results of the nonlinear ARDL model when the explanatory variables are split into two partial sums are given in the right panel in Table 1. In comparison to the ARDL model, the goodness of fit measured as R^2 and adjusted R^2 are higher. However, both the AIC and SIC are higher. Thus, for model selection, we have mixed results for determining which model is more appropriate. As noted before, the ARDL model has a specification error of non-stable coefficient according to the CUSUMSQ test. Our nonlinear ARDL model passes all the diagnostic tests considered in this paper, indicating that the nonlinear ARDL model is more appropriate than the ARDL. While Wald tests cannot reject the long-run symmetries, only positive changes in temperatures and real GDP per capita have significant impacts on arrivals. On the other hand, the negative developments in temperatures and real GDP per capita are not significantly different from zero. More specifically, the estimated long-run coefficient for temperatures L^+_{TEMP} is −0.077 and significant at the five per cent level. However, L^-_{TEMP} is not significantly different from zero. This means that a 1 °C increase in average winter temperatures will lead to an eight per cent decline in arrivals on average, while a decline in temperatures does not lead to a significant increase in arrivals. The result is not surprising since higher temperatures make snow-based winter sport activities less attractive and favour other outdoor activities, thereby reducing tourism demand in winter sport destinations. However, colder winter seasons would not necessarily make snow-based winter sport activities more attractive.

The long-run coefficients L^+_{GDPPC} and L^-_{GDPPC} are 3.5 and 3.1, respectively, but the latter is insignificant at conventional significance levels. This indicates that an increase in real GDP per capita of the visitor countries leads to an increase in arrivals, but a recession is not associated with a decline in arrivals. Interestingly, the long-run coefficients related to currency appreciations and depreciations are not significant, although the long-run coefficient is significant in the symmetric model. Instead, the short-run adjustments become significant and the Wald test identifies asymmetric adjustments. To sum up, the asymmetric reactions of arrivals in terms of magnitude and significances, with respect to increases and decreases of real GDP per capita and temperatures, are a new finding in

Table 1
Determinants of arrivals (ARDL and nonlinear ARDL estimations).

symmetric model (ARDL)			asymmetric model (nonlinear ARDL)		
	coeff.	t-value		coeff.	t-value
α_1	−0.267***	−5.11	α_1	−0.233***	−4.98
L_{GDPPC}	2.396***	21.26	L^{+}_{GDPPC}	3.539***	5.29
			L^{-}_{GDPPC}	3.129	0.67
L_{EX}	0.770**	2.60	L^{+}_{EX}	0.102	0.18
			L^{-}_{EX}	1.562	1.12
L_{TEMP}	−0.046**	−2.26	L^{+}_{TEMP}	−0.077**	−2.19
			L^{-}_{TEMP}	−0.071	−1.13
α_0	−9.318***	−8.30	α_0	11.881**	68.60
β_1	−0.121	−1.05	β_1	−0.117	−1.02
β_2	−0.243**	−2.11	β_2	−0.313***	−2.81
β_0^{GDPPC}	0.841**	2.48	β_0^{GDPPC+}	1.025**	2.23
			β_0^{GDPPC-}	−0.010	0.61
β_0^{EX}	0.161	0.97	β_0^{EX+}	−0.609**	−2.37
			β_1^{EX+}	0.631	1.43
			β_0^{EX-}	0.863**	2.65
β_0^{TEMP}	−0.011**	−2.48	β_0^{TEMP+}	−0.013**	−2.70
			β_0^{TEMP-}	−0.233	−1.31
F_{PSS}	7.55^{aaa}		F_{PSS}	$4.83^{aaa,bbb}$	
$R^2/adj\ R^2$	0.32/0.23		$R^2/adj\ R^2$	0.44/0.27	
AIC/SIC	−2.94/-2.68		*AIC/SIC*	−2.90/-2.42	
χ^2_H	10.22		χ^2_H	14.40	
$\chi^2_{SC}(8)$	5.51		$\chi^2_{SC}(8)$	12.51	
χ^2_{FF}	1.55		χ^2_{FF}	0.24	
χ^2_{normal}	0.64		χ^2_{normal}	2.06	
CUSUM	stable		CUSUM	stable	
CUSUMSQ	not stable		CUSUMSQ	stable	
			W^{GDP}_{LR}	0.01	
			W^{EX}_{LR}	1.61	
			W^{TEMP}_{LR}	0.00	
			W^{GDP}_{SR}	0.01	
			W^{EX}_{SR}	6.53**	
			W^{TEMP}_{SR}	0.00	

1 Notes: Asteriks ***and ** indicate significance at 1% and 5%, respectively. L_{GDP}, L_{EX} and L_{TEMP} denote the long-run elasticities of tourism demand. χ^2_H, $\chi^2_{SC}(8)$, χ^2_{FF} and χ^2_{normal} provide the diagnostic statistics for the Breusch-Pagan's heteroscedasticity, the Breusch-Godfrey's serial correlation with eight lags, the Ramsey's RESET function form tests, the normality, CUSUM and CUSUMSQ stability of parameters tests. F_{PSS} reports the PSS's *F*-statistics for the null hypothesis of no cointegration. ^{aaa} indicates that F_{PSS} is larger than the upper bound of 1%, according to the critical values in Narayan (2005). Our sample period refers to 1962–2015 with 53 observations. The critical values for upper bounds with observations of 50, the case II, restrict intercept and no trend, with $k = 3$, are 5.328, 4.002, and 3.398 corresponding to 1%, 5%, or 10%, respectively, as reported in Narayan (2005). ^{bbb} indicates that F_{PSS} is larger than the upper bound of 1%, according to the critical values with $k = 6$: 3.976, 2.794, and 2.334 corresponding to 1%, 5%, or 10%, respectively, as reported in Narayan (2005). W_{LR} is the Wald test of the long-run symmetry with the null of $L^{+} = L^{-}$. W_{SR} is the Wald test of the null of the short-run symmetry.

the literature. The adjustment coefficient is −0.23, indicating that 23 per cent of the deviation to the long term equilibrium is corrected each year.

Having found a significant but asymmetric relationship between arrivals, GDP, and temperature, the dynamics of temperature coefficients are studied. Here the main hypothesis is that the role of temperature in determining winter tourism demand diminishes over time due to the widespread use of snowmaking and other adaption measures. The choice of the subsample is dictated by the massive investments in snowmaking facilities starting at the beginning of the 1990s after two extremely warm (1988/1989 and 1989/1990) and three snow-poor winter seasons (1987/1988, 1988/1989, and 1989/1990). Between 1994 and 2000, the number of snowmaking facilities (e.g. snow cannons) increased by 12 per cent per year on average (from about 500 to 1100).[9]

The results for the subsample estimations (1960–1991 and 1986–2015) show that the long-run coefficients of temperature decreases and increases significantly change over the two periods (Table 2). Interestingly, while temperature increases have a significant impact on arrivals in the first period (with L^{+}_{TEMP} equal to −0.052), temperature increases are no longer relevant from the 1990s onwards with a long-run coefficient of L^{+}_{TEMP} equal to −0.005. An explanation is the rapid development of snowmaking facilities and improvement in snowmaking technologies, both of which make skiing demand less sensitive to positive temperature shocks.

However, temperature decreases are significant in the second subsample, indicating that tourists prefer colder temperatures given the fact that the average temperature in this period was rather high. The long-run coefficient of L^{-}_{TEMP} is −0.036, indicating that a decrease in temperatures of 1 °C stimulates arrivals by four per cent. Again, the magnitude of negative temperature shocks is significantly lower in absolute terms in the most recent period than in the earlier period (with a temperature coefficient of −0.079). Overall, the finding that the role of temperatures for winter tourism demand significantly declines is consistent with those for the relationship between snow conditions and winter tourism demand (Töglhofer et al., 2011).

Several more robustness checks are conducted (results are not shown here). First, other subsamples are used to check the stability of relationships between temperatures and winter tourism demand. Using the starting year 1988 or 1989 for the second subsample leads to similar long-run coefficients of L^{-}_{TEMP} and of L^{+}_{TEMP}. Second, we use dummy variables for the year 1991 to account for the impact of the German unification. Results with respect to income elasticity are not sensitive to the

[9] Source: Provincia Autonoma di Bolzano-Alto Adige. Ufficio trasporti funiviari, various issues.

Table 2
Sub-sample partial asymmetric models.

1960–1991			1986–2015		
	coeff.	t-value		coeff.	t-value
α_1	−0.419***	−4.19	α_1	−0.649***	−4.59
L_{GDPPC}	2.366***	2.84	L_{GDPPC}	0.463	1.23
LEX	0.222	0.97	L_{EX}	−0.180	−0.91
L^{+}_{TEMP}	−0.052**	−2.07	L^{+}_{TEMP}	−0.005	−0.50
L^{-}_{TEMP}	−0.079*	−1.82	L^{-}_{TEMP}	−0.036***	−4.23
α_0	−9.364	−1.29	α_0	9.307***	94.02
β_0^{GDPPC}	1.420**	2.58	β_0^{GDPPC}	0.526	1.47
β_0^{EX}	0.109	0.49	β_0^{EX}	−0.115	−0.79
			β_1^{EX}	0.377**	2.61
β_0^{TEMP+}	−0.030**	−2.74	β_0^{TEMP+}	−0.003	−0.61
β_0^{TEMP-}	−0.013	−0.91	β_0^{TEMP-}	−0.018***	−2.83
F_{PSS}	5.23^{cc,dd}		F_{PSS}	5.55^{cc,dd}	
R^2/adj R^2	0.34/0.16		R^2/adj R^2	0.47/0.30	
AIC/SIC	−2.51/-2.19		*AIC/SIC*	−3.46/-3.09	
χ^2_H	7.75		χ^2_H	3.09	
$\chi^2_{SC}(4)$	7.75		$\chi^2_{SC}(4)$	4.67	
χ^2_{FF}	3.09*		χ^2_{FF}	0.34	
χ^2_{normal}	2.06		χ^2_{normal}	1.94	
ARCH(2)	0.25		ARCH(2)	0.02	
CUSUM	stable		CUSUM	stable	
CUSUMSQ	stable		CUSUMSQ	stable	
W^{TEMP}_{LR}	0.30		W^{TEMP}_{LR}	16.37***	
W^{TEMP}_{SR}	0.03		W^{TEMP}_{SR}	6.29**	

Notes: Asteriks ***, **, and * indicate significance at 1%, 5%, and 10%, respectively. L_{GDP}, L_{EX} and L_{TEMP} denote the long-run elasticities of tourism demand. χ^2_H, $\chi^2_{SC}(4)$, χ^2_{FF}, χ^2_{normal}, ARCH(2) provide the diagnostic statistics for the Breusch-Pagan's heteroscedasticity, the Breusch-Godfrey's serial correlation with four lags, the Ramsey's RESET function form tests, the normality, ARCH with two lags, CUSUM and CUSUMSQ stability of parameters tests. F_{PSS} reports the PSS's *F*-statistics for the null hypothesis of no cointegration. ^{cc} indicates that F_{PSS} is larger than the upper bound of 5%, in case of $k = 3$, according to the critical values in Narayan (2005). ^{dd} indicates that F_{PSS} is larger than the upper bound of 1%, 5%, or 10%, respectively, in case of $k = 4$, according to Narayan (2005). W_{LR} is the Wald test of the long-run symmetry with the null of $L^{+} = L^{-}$. W_{SR} is the Wald test of the null of the short-run symmetry. The sizes of our subsamples are 30 observations. The critical values for upper bounds with observations of 30, the case II, restrict intercept and no trend, and with $k = 3$ are 5.966, 4.306, and 3.586 corresponding to 1%, 5%, or 10%, respectively, as reported in Narayan (2005). The critical values for upper bounds with observations of 30, the case II, restrict intercept and no trend, and with $k = 4$ are 5.840, 4.223, and 3.560 corresponding to 1%, 5%, or 10%, respectively, as reported in Narayan (2005).

inclusion of the dummy variable (results are available upon request).

5. Conclusions

This study provided additional empirical insights on the relationship between tourism demand and temperatures in the winter season using time-series data for the period 1960–2015. We employ the recently developed nonlinear ARDL cointegration approach to examine the possible long-run and short-run asymmetries of the relationship between temperatures and arrivals in the winter season. A special feature of the model is that it can distinguish between the long-run tourism effects of temperature increases and decreases. The key result of the study is that the impact of positive temperature shocks significantly declined over the sample period. In the period before the widespread usage of snowmaking systems, temperature increases significantly influenced arrivals with a relatively large magnitude. In particular, for the sample period 1962–1991, an increase in temperatures by 1 °C led to a decrease in arrivals by five percent. However, for the recent period from the 1990s onwards, a temperature increase did not lead to a decline in winter tourism demand. The reduction of the sensitivity of winter tourism demand is likely due to the rapid increase in snowmaking facilities from the early 1990s onwards. Another explanation might be that tourists are becoming used to warmer temperatures.

Overall, the nonlinear ARDL shows clear evidence of long-run asymmetry for temperatures and real GDP per capita. In particular, for the latter period, arrivals had a significantly stronger reaction to negative temperature changes than to positive ones in the long-run. However, for the earlier period, positive temperature shocks (unfavourable conditions) had a greater effect on arrivals in absolute terms than negative temperature shocks. Growth of real income led to an increase in the number of arrivals, while a drop in real GDP per capita did not reduce tourism demand.

Our results have several implications for ski lift companies, hospitality industry, destination marketing organisations (DMOs), and government institutions in winter sport regions. First, warm winter seasons are not necessarily a threat to the tourism industry in winter destinations despite the strong increase in temperatures over the last decades. On the whole, the sensitivity of tourism inflows with respect to year-to-year fluctuations in temperatures started to decline from the end of 1980s onwards. However, colder seasons stimulate tourism demand, indicating that winter tourists prefer colder temperatures over warmer temperatures. Second, while warmer seasons do not have an impact on tourism demand, most probably due to the use of snowmaking facilities, this has not come without a cost. Rising temperatures will increase the costs of snowmaking and thereby increase the price of ski lift tickets (Damm et al., 2014). In addition, further investments in snowmaking facilities cannot be compatible with long-term ecological sustainability (De Jong, 2015). It is well known that snow making requires large amounts of water and electricity. In some cases mountain water reserves have to be built or expanded.

This study has several limitations. It is not clear if the relationship between year-to-year fluctuations in temperatures and winter tourism demand reflects a true response to changes in climate factors or rather a short-run weather elasticity (Auffhammer et al., 2013; Rosselló-Nadal, 2014). Furthermore, estimates are based on aggregate data for the whole region in

the winter season. This region not only includes ski resorts and other winter sport destinations but also the capital city and other cities in the valley region where winter sport activities are not relevant. The tourism share of these cities is less than 10 per cent and the exclusion of these regions would be unlikely to have an impact on the estimates. However, it could be the case that the decline of temperature impact on winter tourism demand reflects a change in the preferences of tourists towards destinations in the region that are less dependent on favourable winter temperatures. More detailed data is needed as to whether the declining impact of snow depth is due to the adjustment behaviour of tourists. A second limitation relates to the lack of snow depth data for the sample period.

With respect to future work, it would be preferable to provide estimates for different locations (elevation, size). However, consistent time series on arrivals at the village level over a longer time period are difficult to obtain. Another area of future research would be to account for snow depth. Long-run time series on snow depth are available for the South-western Italian Alps (Terzago et al., 2010). The nonlinear ARDL approach can also be applied to study the link between year-to-year fluctuations and economic outcomes in other industries (for instance, agriculture).

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Appendix

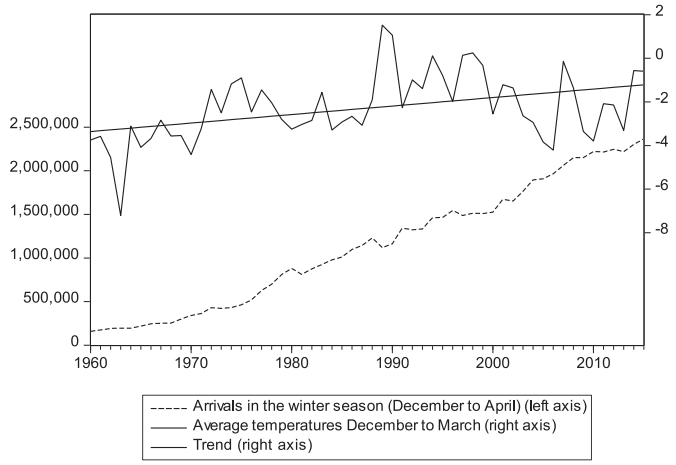

Fig. 1. Evolution of arrivals in the winter season and winter temperatures (month 12 to month 4) at the Sexten weather station (1300 m).
Source: Statistical office south Tyrol. Autonome Provinz Bozen.

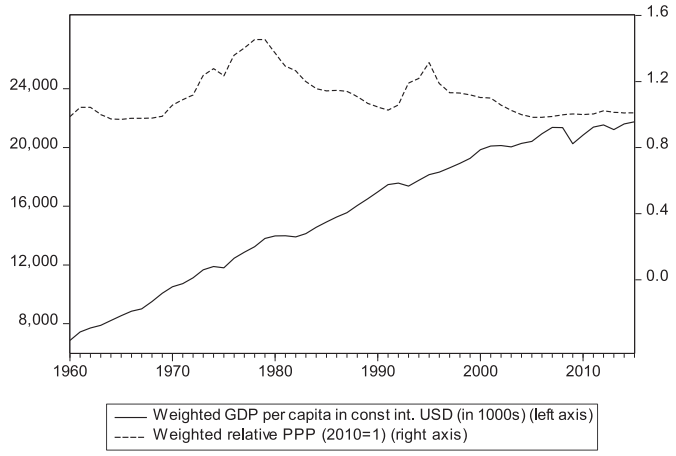

Fig. 2. Evolution of weighted GDP per capita and relative purchasing power parities.
Source: World Bank, Penn world tables.

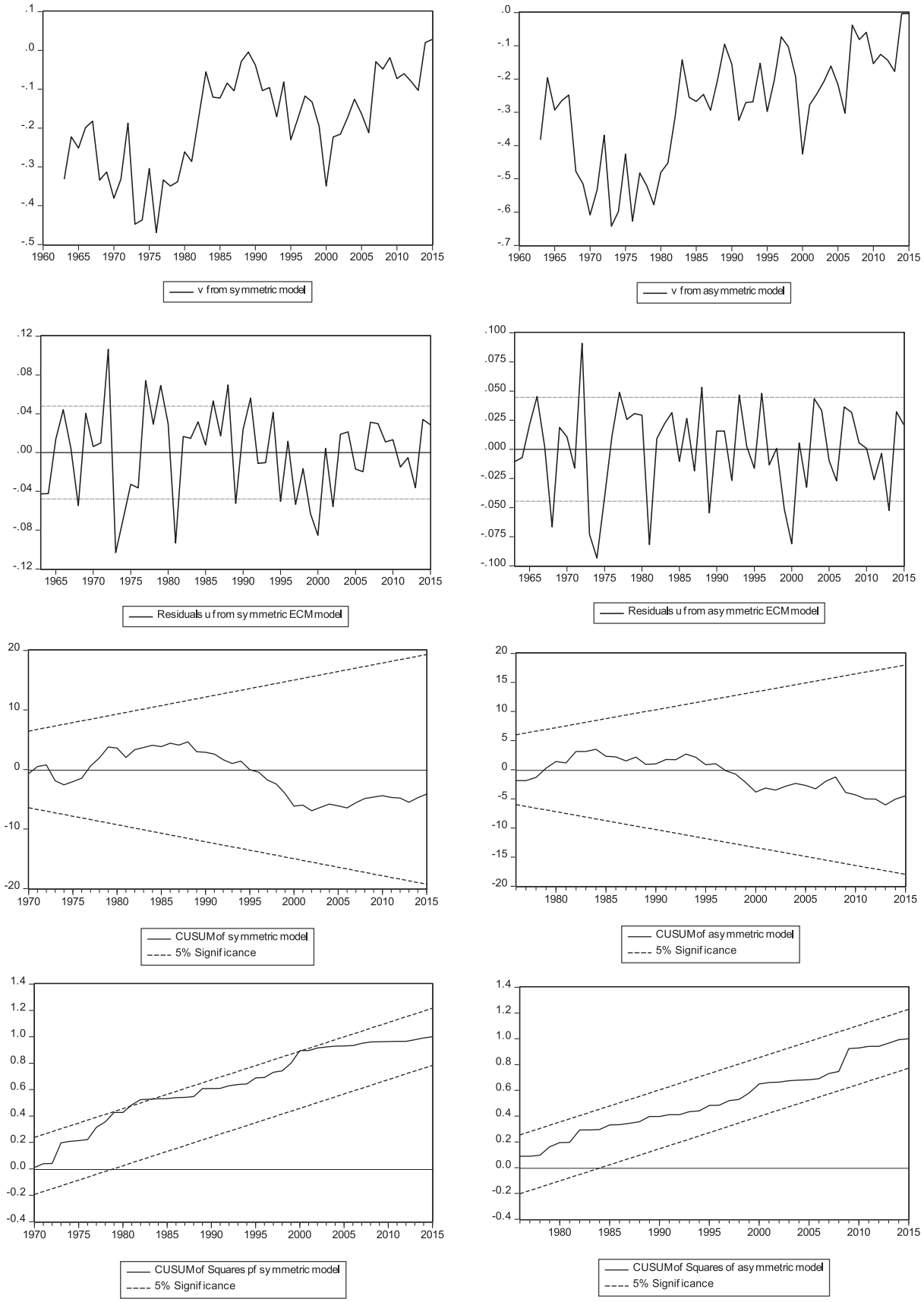
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Fig. 3. Diagnostic tests.

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